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Exchange risk premia in the European monetary system

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In this article, a survey database of exchange rate expectations is employed to examine EMS exchange risk premia. We are able to test a risk premium model directly, i.e. without having to rely on the rational expectations assumption. The results indicate that time-varying risk premia are almost always present and that a (G)ARCH-in-mean specification is often quite succesful in capturing the essential features of the premia.

I. INTRODUCTION

One of the well established empirical regularities in the international financial economics literature is the finding that the forward discount is a biased predictor of the future change in the exchange rate – see the surveys on the efficiency of the foreign exchange market by Hodrick (1987) and Levich (1985). The rejection of forward market efficiency may be attributable to the irrationality of market participants (as suggested, e.g., by Bilson, 1981; Cumby and Obstfeld, 1984 and Longworth, 1981), or to the existence of time-varying risk premia (as suggested by Fama, 1984; Hodrick and Srivastava, 1984; Hsieh, 1984; Wolff, 1987 and others), or to some combination of both of these phenomena. The debate regarding the relative size and variability of the exchange risk premium continues to be an issue of central concern in the financial economics literature. Conditional on the hypothesis that the foreign exchange market is efficient or rational, the existence of time-varying premia has been documented in the literature by Hansen and Hodrick (1980), Frankel (1982), Fama (1984), Hodrick and Srivastava (1984), Hsieh (1984), Korajczyk (1985), and Wolff (1987).

Alternative methodologies to measure time-varying premia have been explored in the literature. First, models that are based strictly on the time series properties of spot and forward exchange rates and asset prices were examined – see the latent variable model of Hansen and Hodrick (1980) with its extensions by Hodrick and Srivastava (1984), Campbell and Clarida (1987), and Giovannini and Jorion (1987). The presence of conditional heteroscedasticity in forecast errors prompted Domowitz and Hakkio (1985) to model a time-varying risk premium using the autoregressive conditional heteroscedasticity (ARCH) framework. Korajczyk (1985) noted that the variability of risk premia can in theory be related to variations in expected real interest rates. A second approach is to employ some measure of market fundamentals in an attempt to test specific theories of the risk premium. Frankel (1982) and Frankel and Engel (1984) examined an asset market equilibrium model based on assets demands derived from a two-period mean-variance maximization problem. A third approach to investigate the possible existence of time-varying risk premia attempts to measure expected depreciation directly using information from surveys – see Frankel and Froot (1987b), Froot and Frankel (1989), and Cavaglia *et al.* (1994), for instance.

This paper modifies the analysis of Domowitz and Hakkio (1985) to be applied to a survey data set of exchange rate expectations covering a wide range of EMS currencies over a different sample period, combining the first and third approaches mentioned above. The principal

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benefit of using such data is that one obtains a direct measure of agents' beliefs, thus allowing for separate testing of an underlying model of exchange rate determination and a hypothesis about expectations, whereas previous work proceeded conditional on the hypothesis that the foreign exchange market is efficient or rational. The survey data set begins in January 1986 and ends in September 1991, when the survey was discontinued, covering a period of US Dollar depreciation (and Deutschmark appreciation) relative to the currencies we review. Cavaglia *et al.* (1994) recently examined a set of survey data of exchange rate expectations, that includes several EMS currencies. Their findings indicate that for EMS exchange rates relative to the Deutschmark variation in the forward discount primarily reflects changes in risk premia rather than changes in expected depreciation and, thus, that the forward discount bias is primarily attributable to significant variation in the risk premium component. As in Domowitz and Hakkio (1985) and Diebold and Pauly (1988), we employ exchange rate data on a monthly basis. However, the results, covering nearly all EMS currencies, provide an interesting complement to previous work that has largely focused on the five most actively traded currencies *vis-à-vis* the US Dollar. Conditional heteroscedasticity appears to be a prominent feature of exchange rate behaviour in the EMS period – see Diebold and Pauly (1988b), and Nieuwland *et al.* (1994), for instance.

This paper is presented in five sections. In Section II, the construction of the exchange rate survey is outlined and summary statistics describing the data are provided. In Section III, the presence of time-varying risk premia is examined as well as the presence of heteroscedastic residuals. The methodology and models employed to capture the time-varying risk premia are explained in Section IV. Section V presents the main empirical results of this study and Section VI contains concluding remarks.

II. THE SURVEY DATA

From 1986 through 1991, Business International Corporation conducted a monthly survey of exchange rate expectations covering five currencies relative to the Deutschmark which are published in its Cross Rates Bulletin. For publication purposes, survey participants are asked a few days prior to month's end to fax three-, six- and twelve-months-ahead expectations of a number of currencies with projections being made from the beginning of the following month. Thus, for instance, the three-, six- and twelve-months-ahead expected French Franc/Deutschmark rate recorded on 27 December 1990 reflect a slightly longer forecast horizon as they represent the expected spot rate on 1 April 1991, 1 June 1991 and 2 January 1992 respectively. The dates on which the surveys were conducted were recorded as well as the spot, three-, six-, and twelve-month-ahead forward rates recorded on that particular day. Care has been exercised throughout the empirical analysis to ensure that conditional expectations are computed on the basis of the proper information set.

The thirty-odd participants of the survey are treasurers of multinationals and private banks residing in four of the world's continents. Although not all participants provide their views regarding a particular currency, the response rate is at worst 60%. The *Cross Rates Bulletin* reports the geometric mean forecast of the responses received, thus minimizing the effect of extreme forecasts. Unfortunately disaggregated survey respondent data are not available, although the standard deviation of the respondents' expectation is reported.

Conditional on market efficiency and rational expectations, the forward exchange rate is equal to the expected future spot rate plus a risk premium. The use of survey data allows the direct measurement of a risk premium from the decomposition of the forward discount into its two components – expected depreciation and the risk premium:

$$ {}_tF_{t+k} - S_t = (E_tS_{t+k} - S_t) + P_t^k \tag{1}$$

Here S_t is defined as the natural logarithm of the spot exchange rate at time t, E_tS_{t+k} is defined as the expected logarithm of the spot exchange rate at time $t+k$ formed at time t and ${}_tF_{t+k}$ is defined as the natural logarithm of the forward rate at time t for delivery at time $t+k$ and P_t^k is the associated risk premium. The left-hand side of Equation 1 is the forward discount, and the right-hand side is the expected rate of depreciation of the home currency relative to the foreign currency (Deutschmark – the exchange rates are expressed as units of home currency per unit of foreign currency) plus the risk premium.

Because the exchange rate expectations from the survey are direct estimates they do not require us to assume any particular model of expected depreciation or of the risk premium. To give Equation 1 economic content, a model of international asset pricing that describes the determination of P_t^k is required. Equilibrium models of international asset pricing that provide us with such descriptions are presented, for instance, in Roll and Solnik (1977), Hodrick (1981), Stulz (1981), Adler and Dumas (1983), and Hodrick and Srivastava (1984).

Tables 1(a) and (b) provide summary statistics for the expected exchange rate depreciation and forward discount across forecast horizon and across currencies. The summary statistics for the risk premium across horizon and across currencies are reported in Table 1(c).

For the period analysed (1 January 1986 through 1 September 1991) the standard deviations of both the expected depreciation and forward discount across the 3, 6, and 12 month horizons are generally larger for the Italian Lira relative to the Deutschmark than the corre-

Table 1(a). *Summary of statistics of expected depreciation:* $E_tS_{t+k} - S_t$: *1 January 1986 through 1 September 1991*

	BF/DM	DG/DM	FF/DM	IL/DM	SP/DM
			3 Months		
Mean	0.0018	0.0028	0.0048	0.0067	0.0121
St. dev	0.0107	0.0098	0.0063	0.0257	0.0147
T-test	1.41	2.34	6.41	2.16	6.84
Skewness	2.15	0.11	0.26	−2.52	0.65
Kurtosis	11.14	10.87	3.48	18.36	4.84
BJ-test	243.42***	178.16***	1.45	750.80***	14.60***
KS-1	52.94***	0.14	0.78	72.86***	4.83**
KS-2	190.49***	178.02***	0.67	677.93***	9.77***
			6 Months		
Mean	0.0046	−0.0010	0.0077	0.0122	0.0198
St. dev.	0.0114	0.0078	0.0125	0.0242	0.0202
T-test	3.35	−1.08	5.11	4.19	8.14
Skewness	1.13	−1.86	0.20	−2.90	0.83
Kurtosis	5.60	11.15	2.76	19.47	4.57
BJ-test	34.09***	230.50***	0.65	876.49***	15.07***
KS-1	14.63***	39.58***	0.48	96.52***	7.97***
KS-2	19.46***	190.92***	0.17	777.97***	7.10***
			12 Months		
Mean	0.0060	−0.0024	0.0122	0.0168	0.0252
St. dev	0.0147	0.0102	0.0144	0.0306	0.0196
T-test	3.43	−1.93	7.03	4.54	10.67
Skewness	0.79	−0.31	0.16	−1.50	0.60
Kurtosis	4.20	5.07	2.65	10.17	3.32
BJ-test	11.39***	13.44***	0.65	173.64***	4.37**
KS-1	7.26***	1.11	0.30	28.89***	4.08**
KS-2	4.13**	12.33***	0.35	147.74***	0.29

Notes: BF = Belgian Franc; DG = Dutch Guilder; DM = Deutschmark; FF = French Franc; IL = Italian Lira; SP = Spanish Peseta. The BJ-test denotes the Bera–Jarque test for normality; KS-1 and KS-2 pertain to the Kiefer–Salmon normality test for skewness and kurtosis, respectively; * (**) [***] denotes rejection at the 10% (5%) [1%] level of the normality hypotheses.

Table 1(b). *Summary of statistics of forward discount:* ${}_tF_{t+k} - S_t$: *1 January 1986 through 1 September 1991*

	BF/DM	DG/DM	FF/DM	IL/DM	SP/DM
			3 Months		
Mean	0.0062	0.0017	0.0080	0.0159	0.0170
St. dev	0.0084	0.0053	0.0050	0.0110	0.0088
T-test	6.09	2.62	13.48	12.01	15.94
Skewness	0.72	−4.69	0.79	4.67	−0.25
Kurtosis	12.57	39.45	3.95	32.08	2.73
BJ-test	269.19***	4072***	9.89***	2681***	0.92
KS-1	6.01**	253.68***	7.31***	250.37***	0.71
KS-2	263.17***	3818***	2.58	2430***	0.22
			6 Months		
Mean	0.0112	0.0027	0.0151	0.0312	0.0314
St. dev.	0.0114	0.0059	0.0092	0.0270	0.0167
T-test	8.11	3.78	3.61	9.60	15.63
Skewness	2.19	−3.65	0.66	6.11	−0.65
Kurtosis	18.97	28.01	3.27	46.06	2.94
BJ-test	788.23***	1951***	5.37**	5760***	4.84**
KS-1	54.94***	153.03***	5.15**	429.34***	4.83**
KS-2	733.29***	1798***	0.22	5331***	0.01
			12 Months		
Mean	0.0209	0.0064	0.0285	0.0584	0.0587
St. dev	0.0170	0.0137	0.0168	0.0431	0.0297
T-test	10.21	3.88	14.11	11.25	16.45
Skewness	1.12	4.57	0.41	5.70	−0.76
Kurtosis	11.04	35.94	2.56	42.36	3.10
BJ-test	200.43***	3360***	2.46	4828***	6.61**
KS-1	14.53***	240.72***	1.89	374.14***	6.58**
KS-2	185.90***	3119***	0.57	4454***	0.03

Notes: BF = Belgian Franc; DG = Dutch Guilder; DM = Deutschmark; FF = French Franc; IL = Italian Lira; SP = Spanish Peseta. The BJ-test denotes the Bera–Jarque test for normality; KS-1 and KS-2 pertain to the Kiefer–Salmon normality test for skewness and kurtosis, respectively; * (**) [***] denotes rejection at the 10% (5%) [1%] level of the normality hypotheses.

sponding estimates of the other EMS exchange rates involving the Deutschmark. The provisions of the EMS at the time allowed participating countries to maintain their exchange rates within bilateral limits of $\pm 2.25\%$ ($\pm 6\%$ for Italy and, since June 1989, Spain). Comparing Table 1(a) and (b), one notes that in general the expected rates of depreciation and the forward discount are of the same sign. Thus the currencies that were expected to depreciate were at a forward discount. This confirms the results of Frankel and Froot (1987a, b). Table 1(c) suggests the presence of time-varying risk premia, implying that domestic and foreign assets are imperfect substitutes. The numbers differ from summary statistics reported by Frankel and Froot (1987a, b), which demonstrate surprisingly large exchange risk premia in a number of cases.

In order to assess the distributional properties of the expected depreciation, forward discount, and risk premia series, the Bera–Jarque (1982) normality test and the Kiefer–Salmon (1983) Lagrange multiplier normality tests are reported in Table 1(a), (b), and (c). The former is a joint test using both skewness and kurtosis and the latter are a LM tests for normal skewness (KS-1) and normal kurtosis (KS-2), respectively. The Bera–Jarque test is asymptotically $\chi^2(2)$ distributed and the Kiefer–Salmon tests $\chi^2(1)$. Overall, the evidence presented suggests a fairly consistent rejection of the normality hypotheses. Failure to reject the null occurs in only seven out of 45 cases. Thus, in spite of the notion that leptokurtic unconditional densities of ARCH processes approach normality by temporal aggregation – see e.g. Diebold (1988b) – it appears that the monthly series used here may be characterized as highly leptokurtic. This is in line with Koedijk *et al.* (1990) who find that for EMS exchange rates ARCH effects become less important in time aggregation whereas fat tails remain important.

Table 1(c). *Summary statistics of risk premium:* P_t^k: *1 January 1986 through 1 September 1991*

	BF/DM	DG/DM	FF/DM	IL/DM	SP/DM
			3 Months		
Mean	0.0044	−0.0011	0.0032	0.0092	0.0049
St. dev	0.0125	0.0085	0.0052	0.0250	0.0144
T-test	2.91	−1.08	5.16	3.07	2.80
Skewness	−0.63	−1.16	0.49	2.96	0.27
Kurtosis	8.37	6.90	4.68	16.21	5.88
BJ-test	87.64***	59.12***	10.82***	602.98***	24.74***
KS-1	4.59**	15.37***	2.71*	100.94***	0.83
KS-2	83.05***	43.75***	8.11***	502.04***	23.91***
			6 Months		
Mean	0.0065	0.0037	0.0074	0.0190	0.0116
St. dev.	0.0133	0.0067	0.0086	0.0329	0.0221
T-test	4.09	4.57	7.08	4.79	4.38
Skewness	1.84	1.01	0.15	5.18	−0.58
Kurtosis	12.94	5.11	3.02	33.01	6.39
BJ-test	323.18***	24.43**	0.27	2898***	36.88***
KS-1	38.99***	11.65***	0.27	308.86***	3.92**
KS-2	284.19***	12.78***	0.00	2590***	32.96***
			12 Months		
Mean	0.0149	0.0088	0.0163	0.0417	0.0336
St. dev	0.0191	0.0162	0.0121	0.0490	0.0296
T-test	6.46	4.48	11.19	7.06	9.42
Skewness	1.82	4.17	1.03	5.19	−0.16
Kurtosis	11.22	28.84	4.81	37.55	2.64
BJ-test	232.30***	2119***	21.73***	37.41***	0.67
KS-1	38.06***	199.83***	12.27	309.54***	0.30
KS-2	194.23***	1919***	9.47***	3432***	0.37

Notes: BF = Belgian Franc; DG = Dutch Guilder; DM = Deutschmark; FF = French Franc; IL = Italian Lira; SP = Spanish Peseta. The BJ-test denotes the Bera–Jarque test for normality; KS-1 and KS-2 pertain to the Kiefer–Salmon normality test for skewness and kurtosis, respectively; * (**) [***] denotes rejection at the 10% (5%) [1%] level of the normality hypotheses.

III. TIME-VARYING EXCHANGE RISK PREMIA

Survey expectations data can be exploited to decompose the forward discount bias into portions attributable to irrational behaviour of economic agents or to the existence of time-varying risk premia – see Frankel and Froot (1987b), Froot and Frankel (1989), and Cavaglia *et al.* (1993, 1994), for instance. In order to test whether the existence of time-varying risk premia is the economically important reason for rejection of forward market efficiency, the following equation may be fitted by ordinary least squares:

$$E_tS_{t+k} - S_t = \alpha + \beta({}_tF_{t+k} - S_t) + \varepsilon_t \tag{2}$$

where ε_t is a random error term. The null hypothesis of perfect substitutability implies that $\alpha = 0$ and $\beta = 1$. The degree to which changes in the forward discount reflect changes in the risk premium can be inferred from a regression of expected depreciation on the forward discount (Equation 2). Under the hypothesis that the correlation of the risk premium with the forward discount is zero (no time-varying risk premia), β will equal one. Cavaglia, *et al.* (1993, 1994) and Cavaglia and Wolff (1993) present evidence for the dataset at hand, based on heteroscedasticity-consistent test statistics, that perfect substitutability is almost always rejected and that the hypothesis $\beta = 1$ is also fairly consistently rejected. Thus, as in most models in which sterilized foreign exchange intervention is effective, variation in the forward discount for EMS currencies reflect a statistically significant degree of variation in the risk premium component.

In order to test for the presence of heteroscedastic OLS residuals ε_t from Equation 2, two different approaches are employed below. First the Lagrange multiplier (LM) tests for autoregressive conditional heteroscedasticity – see Breusch and Pagan (1979) – are performed, and secondly a nonparametric test based on finite-state homogeneous Markov chains – see Gregory (1989) – is applied. Using Monte Carlo analysis, Gregory (1989) concludes that under other distributions than the Normal the LM test is biased towards the null hypothesis of no ARCH, and that the Markov Chain test is superior to the LM test in terms of better finite sample properties. Both tests only require estimation under the null hypothesis of no heteroscedasticity and are appropriate under all distributional assumptions.

The results of the LM and Markov chain tests for the presence of heteroscedasticity are given in Table 2. Overall, the evidence presented suggests a weak rejection of the hypothesis of no heteroscedasticity. However, it is interesting to note that the results for the Belgian Franc and the Italian Lira at the 6 and 12 month horizons provide a strong rejection of the null hypothesis. Thus, although ARCH effects tend to weaken with less frequently sampled data, in several cases the EMS exchange rates at the 3, 6, and 12 month horizon still display significant ARCH effects. The evidence presented contrasts with the results of Domowitz and Hakkio (1985), who found no significant ARCH effects for a different set of currencies, except in the case of Japan.

IV. MODELLING TIME VARYING RISK PREMIA: METHODOLOGY

Hodrick's (1987) and Levich's (1985) reviews of the literature on the efficiency of foreign exchange market suggest that there is overwhelming evidence in favour of the view that forward rates are biased predictors of future spot rates. For the EMS currencies examined, rejection is generally attributed to the presence of a significant time-varying risk premium. A number of theoretical models have been put forward which generate risk premia in

Table 2. *Heteroscedasticity tests of OLS residuals:* $E_tS^{-st}_{t+k} = \alpha + \beta({}_tF_{t+k} - S_t) + e_t$*: 1 January 1986 through 1 September 1991*

	BF/DM	DG/DM	FF/DM	IL/DM	SP/DM
			3 Months		
LM(1)	0.25	1.27	3.29*	1.13	14.01***
LM(2)	0.48	1.44	6.81	4.76	3.52
LM(5)	2.83	3.42	10.13	9.11	29.81***
LRIM1	0.02	5.46**	0.60	0.13	0.38
LRIM2	0.16	6.61*	1.15	3.10	0.39
			6 Months		
LM(1)	20.77**	0.04	1.65	0.27	0.60
LM(2)	30.56***	0.06	2.66	9.22***	1.40
LM(5)	4.90	5.28	5.21	26.05**	4.19
LRIM1	3.38*	0.02	5.46**	3.38*	1.22
LRIM2	4.27	0.40	7.47*	4.35	2.36
			12 Months		
LM(1)	14.37***	0.13	1.28	0.03	0.11
LM(2)	10.86***	0.22	1.29	10.31***	2.22
LM(5)	3.36	1.91	2.83	13.08***	6.63
LRIM1	3.38*	0.37	0.38	8.05***	0.02
LRIM2	4.06	2.23	0.55	9.58**	2.96

Notes: LM(p) test is estimated by a regression of squared OLS residuals (Equation 2) on a constant and p lags, and is asymptotically Chi-square (p) distributed. LRIM1 is a likelihood ratio test of independence against a first order Markov Chain, and is distributed chi-square(1); LRIM2 is a likelihood ratio test of independence against a second order Markov Chain, and is distributed chie-square(3). * (**) [***] denotes significance at the 10% (5%) [1%] level.

foreign exchange markets, examples are Hodrick and Srivastava (1984), Domowitz and Hakkio (1985), Diebold and Pauly (1988a) and Kaminsky and Peruga (1990). Most of these theories share Lucas' (1982) model for the international economy as a starting point. Although this dynamic general equilibrium model provides useful insights into the possible structure of risk premia in the forward foreign exchange markets, direct tests of this model are impossible without further restrictions. This is due to the general structure of the model. The second common denominator in these models is that in general the risk premium depends on the conditional probability distribution of the future spot rate, which may lead to a time-varying risk premium, if this distribution is time-varying. Empirically, many specifications for such a risk premium have been employed which depend on the conditional variance of the spot rate. Nevertheless, in their review Bollerslev *et al.* (1992) note that: 'A satisfactory model for the time varying risk premium in the forward foreign exchange market has yet to be formulated'.

In this paper we adopt a different approach which is inspired by the availability of survey data. We do not assume rational expectations, nor do we have to rely on estimation methods using unobserved variables, as in Hodrick and Srivastava (1984, 1986). Conditional on the hypothesis that the foreign market is efficient or rational, the modelling of time varying risk premia has been explored, among others, by Domowitz and Hakkio (1985), and Diebold and Pauly (1988a). Based on the utility optimizing models of Lucas (1982), Domowitz and Hakkio (1985) present an intertemporal asset pricing model in which the risk premium is a function of the conditional variances of the domestic and foreign money supplies. The methodologies used in these papers usually involve measurement of time-varying risk premia conditional on the hypothesis that exchange rate forecasts are rational. Conclusions about the behaviour of premia in the pricing of forward foreign exchange are conditional on rational expectations formation by economic agents. Since the results of previous research overwhelmingly favour the conclusion that economic agents exhibit irrational behaviour (see Frankel and Froot, 1987a and Cavaglia *et al.*, 1993, 1994), we propose an alternative approach to measure premia based on our survey data. The survey data allow the direct measurement of risk premia from the decomposition of the forward discount in Equation 1, thereby avoiding the rational expectations hypothesis.

In our analysis above, several currencies display significant ARCH effects. However, as we do not specify a general equilibrium model we do not know the true structure of the covariance matrix and to what variables it is related. A (G)ARCH model is an acceptable alternative because it can be interpreted as a reduced form of a more complicated dynamic structure for the time-varying conditional second order moments. The ARCH-M model developed by Engle *et al.* (1987) which was proposed in this context by Domowitz and Hakkio (1985), can be used in addressing questions regarding the risk-return tradeoff in a time series context. The ARCH-in-mean model extends the ARCH model to allow the conditional variance to affect the conditional mean directly. The Domowitz and Hakkio model is given by the following equations:

$$E_tS_{t+k} - S_t = RP^k_t + \beta_1({}_tF_{t+k} - S_t) + \varepsilon_t \tag{3}$$

$$RP^k_t = \beta_0 + \theta h_t \tag{4}$$

$$\varepsilon_t | I_{t-1} \sim N(0, h^2_t) \tag{5}$$

$$h^2_t = \alpha_0 + \alpha_1 \varepsilon_{t-1}{}^2 \tag{6}$$

where I_t represents the set of available information at time t. The risk premium, RP^k_t, depends directly on the conditional variance of ε_t which is denoted h^2_t. The conditional variance of the expected rate of depreciation given time t information is postulated to depend on the realizations of the squared error terms in the previous months. A generalization proposed by Bollerslev (1986) is the GARCH model. For the first-order GARCH-in-mean model the conditional variance becomes:

$$h^2_t = \alpha_0 + \alpha_1 \varepsilon_{t-1}{}^2 + \gamma_1 h_{t-1}{}^2 \tag{7}$$

The degree of persistence in variance is determined by the magnitude of the parameters of the conditional variance Equations 6 and 7 and nonnegativity constraints are imposed on these parameters. In the context of the GARCH model, we restrict our attention to a GARCH (1,1) specification since it has been shown to be a parsimonious representation of conditional variance that adequately fits many economic and financial time series – see e.g. Bollerslev (1986).

The estimation of the econometric model described above using maximum likelihood methods is not as straightforward as may seem at first sight. First, we assume conditional normality without knowing what the true conditional distribution really is. This leaves the model subject to distributional misspecification. Weiss (1986) has shown that the quasi-maximum likelihood parameter estimates are still consistent and asymptotically normal but with a modified asymptotic covariance matrix, which is robust to departures from normality. Furthermore, the information matrix is not block diagonal in the present framework, as a function of the conditional variance enters the mean equation. This means that we cannot use the scoring algorithm to obtain the maximum likelihood estimates (see Engle, 1982). Instead, we have to rely on numerical maximization of the likelihood function. We employ the Berndt *et al.* (BHHH) (1974) algorithm with numerical derivatives to obtain maximum likelihood estimates. The robust covariance matrix is calculated by pre- and post-multiplying the inverse of the BHHH covariance matrix by the inverse of the estimated information matrix. The estimation results and diagnostics of both ARCH- and GARCH-in-mean models are presented in the next section.

V. THE EMPIRICAL RESULTS

Maximum likelihood estimates of the parameters and their heteroscedasticity-consistent asymptotic standard errors are reported Tables 3 and 4 for the two models described

Table 3. *ARCH-in-mean models: 1 January 1986 through 1 September 1991*

	β_0	β_1	θ	$\alpha_0(.10^2)$	α_1	L.L.
			3 Months			
BF/DM	−0.0128	0.2632***	1.3091	0.0067***	0.5003***	218.54
	(0.0151)	(0.1003)	(1.4563)	(0.0019)	(0.2517)	
DG/DM	−0.0018	0.8339	0.3735	0.0069***	0.0006	230.76
	(0.0365)	(0.5004)	(4.3029)	(0.0024)	(1.1530)	
FF/DM	0.0125***	0.7731	−2.8027***	0.0023***	0.0515	268.48
	(0.0045)	(0.1581)	(1.0187)	(0.0005)	(0.0376)	
IL/DM	−0.0176	0.8382	0.5236*	0.0343	1.2005	159.19
	(0.0159)	(0.8929)	(0.3124)	(0.0244)	(1.1738)	
SP/DM	−0.0146	0.4874***	1.3919*	0.0149***	0.1423	201.47
	(0.0136)	(0.1730)	(0.8132)	(0.0052)	(0.1242)	
			6 Months			
BF/DM	−0.0022	0.5399*	0.1091	0.0061**	0.6493	217.83
	(0.0025)	(0.2780)	(0.2699)	(0.0025)	(0.6348)	
DG/DM	0.0101	0.7170	−2.0323	0.0036***	0.1341	251.15
	(0.0149)	(0.2551)	(2.4406)	(0.0011)	(0.3673)	
FF/DM	0.0053	0.9842	−1.3584	0.0065***	0.0950	231.15
	(0.0503)	(0.2061)	(6.0928)	(0.0021)	(0.2568)	
IL/DM	0.0001	0.3432***	0.4893*	0.0088***	1.1044***	182.84
	(0.0088)	(0.1822)	(0.2444)	(0.0018)	(0.2727)	
SP/DM	0.0085	0.2105***	0.2588	0.0155**	0.8533	179.06
	(0.0074)	(0.1333)	(0.3319)	(0.0081)	(0.6892)	
			12 Months			
BF/DM	−0.0076	0.0921***	0.8288***	0.0080***	0.7280**	203.24
	(0.0058)	(0.1101)	(0.2939)	(0.0039)	(0.3534)	
DG/DM	0.0067***	−0.0349***	0.0220	0.0040***	1.0041**	222.25
	(0.0013)	(0.0915)	(0.1232)	(0.0014)	(0.4549)	
FF/DM	−0.0266***	0.5944***	2.1682***	0.0092***	0.0646	220.24
	(0.0070)	(0.0754)	(0.8023)	(0.0017)	(0.0603)	
IL/DM	0.0079	0.1382***	0.2918***	0.0098***	1.3161***	165.88
	(0.0096)	(0.0992)	(0.0929)	(0.0034)	(0.3596)	
SP/DM	−0.0171	0.2195***	1.5984	0.0335***	0.0043	177.98
	(0.0211)	(0.0618)	(1.1739)	(0.0071)	(0.0469)	

Notes: The heteroscedasticity-consistent standard errors of the coefficients are given in parentheses; * (**) [***] denotes significance at the 10% (5%) [1%] level for the hypotheses $\beta_0 = 0$, $\beta_1 = 1$, $\theta = 0$, $\alpha_0 = 0$ or $\alpha_1 = 0$, respectively. L.L. denotes the log-likelihood values.

Table 4. *GARCH-in-mean models: 1 January 1986 through 1 September 1991*

	β_0	β_1	θ	$\alpha_0(\cdot 10^2)$	α_1	γ_1	L.L.
				3 Months			
BF/DM	−0.0135	0.2122***	1.3406***	0.0042	0.4337	0.2457	219.30
	(0.0051)	(0.1064)	(0.5486)	(0.0032)	(0.5036)	(0.3598)	
DG/DM	0.0074**	0.9548	−0.7371*	0.0003	0.000**	0.9499***	232.46
	(0.0032)	(0.2771)	(0.3921)	(0.0003)	(0.0000)	(0.0507)	
FF/DM	0.0020	0.7577*	−0.6475	0.0023***	0.0790	0.0000	268.58
	(0.0066)	(0.1673)	(1.4011)	(0.0005)	(0.0700)	(0.0000)	
IL/DM	−0.0222	2.4866***	−0.8173***	0.0018	0.7323***	0.4779***	181.67
	(0.0063)	(0.4408)	(0.2054)	(0.0013)	(0.2003)	(0.1323)	
SP/DM	−0.0090**	0.2819***	1.4544***	0.0005*	0.0000	0.0233***	211.45
	(0.0037)	(0.1798)	(0.2920)	(0.0003)	(0.0000)	(0.0275)	
				6 Months			
BF/DM	−0.0177***	0.0810***	2.2182***	0.0011	0.1588**	0.6898***	228.25
	(0.0025)	(0.0371)	(0.2991)	(0.0007)	(0.0714)	(0.0566)	
DG/DM	0.0083***	0.7614*	−1.7600***	0.0008	0.1693*	0.6315***	253.87
	(0.0027)	(0.1445)	(0.5177)	(0.0005)	(0.1032)	(0.1852)	
FF/DM	0.0065	0.9514	−1.4950	0.0060**	0.0899	0.0753	231.15
	(0.0189)	(0.1770)	(2.3751)	(0.0027)	(0.1302)	(0.2780)	
IL/DM	0.0031	0.3781***	0.2564	0.0003	0.4060***	0.7236***	186.74
	(0.0130)	(0.2128)	(0.5179)	(0.0010)	(0.1246)	(0.0684)	
SP/DM	(0.0085)	0.2105***	0.2587	0.0155**	0.8536	0.0000	179.06
	(0.0085)	(0.1383)	(0.3474)	(0.0073)	(0.7672)	(0.0885)	
				12 Months			
BF/DM	−0.0076	0.0920***	0.8288***	0.0080**	0.7281**	0.0000	203.24
	(0.0058)	(0.1101)	(0.2940)	(0.0039)	(0.3534)	(0.0000)	
DG/DM	0.0068	0.0395***	−0.7682	0.0019	0.4264	0.4798	244.15
	(0.0071)	(0.2347)	(0.9379)	(0.0013)	(0.6647)	(0.4368)	
FF/DM	−0.0204**	0.5969***	1.6435	0.0041	0.0136	0.5150	221.84
	(0.0114)	(0.1324)	(1.2154)	(0.0097)	(0.3429)	(1.3983)	
IL/DM	0.0073	0.1542***	0.2353	0.0038	0.7959***	0.3568***	167.69
	(0.0076)	(0.0873)	(0.1705)	(0.0026)	(0.2534)	(0.1520)	
SP/DM	0.0215***	0.2653***	−0.6997***	0.0000	0.0000	0.9863***	180.19
	(0.0044)	(0.0655)	(0.2117)	(0.0000)	(0.0000)	(0.0051)	

Notes: Heteroscedasticity-consistent standard errors of the coefficients are given in parentheses; * (**) [***] denotes significance at the 10% (5%) [1%] level for the hypotheses $\beta_0 = 0$, $\beta_1 = 1$, $\theta = 0$, $\alpha_0 = 0$ or $\alpha_1 = 0$, respectively. L.L. denotes the log-likelihood values.

in the previous section. All calculations were performed using the software package Gauss.

Table 3 reports the ARCH-in-mean estimation results for each currency and for each forecast horizon ($k = 3, 6$ and 12 months). Rejection of the hypothesis $\beta_1 = 1$ (no time-varying risk premia) was obtained in 10 out of 15 cases, thus corroborating our earlier results, which demonstrate significant time-varying risk premia. In a number of cases, the results provide evidence of both θ and α_1 being insignificantly different from zero. Rejection of the hypothesis $\theta = 0$ was obtained in 7 out of 15 cases, whereas rejection of $\alpha_1 = 0$ was obtained in 5 cases. The Italian Lira/Deutschmark exchange rate appears to be integrated-in-variance (see Engle and Bollerslev, 1986), a condition analogous to a unit root in the conditional mean. At the 3, 6 and 12 month forecast horizon, the estimated α_1 coefficient is greater than one which implies that the unconditional distribution of the expected depreciation is extremely fat tailed with an infinite variance (see also Table 1(a)).

The GARCH-in-mean estimation results are reported in Table 4. The results provide a fairly consistent rejection of the hypothesis $\beta_1 = 1$ (no time-varying risk premia), suggesting significant variation in the risk premium. The estimated α_1 and γ_1 coefficients are statistically significant in many cases, thus supporting the GARCH specification. Moreover, a number of the estimated models result in statistically significant θ coefficients, suggesting that premia for EMS exchange rates relative to the Deutschmark follow GARCH (1,1) processes. In the case of the Italian lira/Deutschmark exchange rate, the coefficient estimates of $\alpha_1 + \gamma_1$ are greater than one, indicating high persistence in the volatility shocks, or IGARCH behaviour (see Engle and Bollerslev, 1986). As conjectured by Diebold (1988), and Lamoureux and Lastrapes (1990), this may

Table 5. *Generalized likelihood ratio test statistics: ARCH-in-mean model against GARCH-in-mean model: 1 January 1986 through 1 September 1991*

	3 months	6 months	12 months
BF/DM	1.52	20.84***	0.00
	(0.218)	(0.000)	(1.000)
DG/DM	3.40*	5.44**	3.80*
	(0.065)	(0.020)	(0.051)
FF/DM	0.20	0.04	3.20*
	(0.655)	(0.841)	(0.074)
IL/DM	44.96***	7.80***	3.62*
	(0.000)	(0.005)	(0.057)
SP/DM	19.96***	0.00	4.42**
	(0.000)	(1.000)	(0.035)

Notes: *P*-values are given in parentheses; * (**) [***] denotes rejection at the 10% (5%) [1%] level. Significance is assessed by comparison with the $\chi^2(1)$.

be the result of shifts in monetary regimes which affect the level of the unconditional variances. Lastrapes (1989) finds that persistence of exchange rate volatility decreases when regime shifts are accounted for, thus diminishing the likelihood of finding integrated-in-variance processes.

Given the above results, it is interesting to compare the relative fit of both models. We employ generalized likelihood ratio tests to compare nested models. Such nested models can be tested using the generalized likelihood ratio $\Lambda = \sup_{\phi\in\Phi} L(\phi;\mathbf{x})/\sup_{\phi\in\Omega} L(\phi;\mathbf{x})$ of the maximized likelihood values under the null, Φ, and under the encompassing parameter space, Ω, which also includes the alternative hypothesis. Here, $L(.;.)$ is the likelihood function, ϕ is the parameter vector and $\mathbf{x}$ is the relevant set of observations. Under the null Φ, the statistic $-2\ln\Lambda$ has a χ^2 distribution with degrees of freedom equal to the difference in the number of parameters of the two models.

Table 5 presents the generalized likelihood ratio tests to compare the relative fit of the two models. Many of the *p*-values associated with the chi-square statistics are close to zero. Thus, the generalized likelihood ratio tests in these cases reject the simpler (ARCH-in-mean) model in favour of the more complicated (GARCH-in-mean) model. In the case of the BF/DM exchange rate at the 12 month horizon and the SP/DM exchange rate at the 6 month horizon, the *p*-values associated with the chi-square statistics are equal to one, which indicates strong support for the ARCH specification. In order to determine the adequacy of the statistical specification, the models are subjected to diagnostic checks on the standardized residuals:

$$z_t = \varepsilon_t^*/h_t^* \tag{8}$$

where ε_t^* is the fitted residual from Equation 3 and h_t^* is the estimated conditional variance from Equations 6 and 7. From Jensen's inequality it follows that the standardized residuals, z_t, should demonstrate less absolute skewness and should be thinner tailed than their unconditional raw data counterparts. Any strong violation of this rule should be regarded as evidence of model misspecification – see Hsieh (1989).

The diagnostics for ARCH- and GARCH-in-mean models are presented in Tables 6 and 7. Overall, the evidence presented suggest a less consistent rejection of the normality hypotheses as compared with the results of Table 1(a), (b), and (c). For the standardized ARCH-in-mean residuals, rejection occurs in 10 out of 15 cases, whereas for the standardized GARCH-in-mean residuals, rejection was obtained in only seven cases. In addition, we find that in most cases the estimated statistics – the BJ-test, KS-1, and KS-2 – are smaller than those reported by Table 1(a), (b), and (c), thus supporting our model specifications. In particular the GARCH-in-mean model is quite succesful at removing excess kurtosis and skewness in a number of cases. In order to test for remaining heteroscedasticity, a residuals-based test of the models may be carried out by regressing $(\varepsilon_t^2 - h_t^2)/h_t^2$, calculated in the basis of fitted values, on fitted values of $1/h_t^2$ and on one to five lags of the dependent variable. The results are reported under LM(1) and LM(5), and are chi-square distributed with one and five degrees of freedom, respectively. For the

Table 6. *Diagnostics of ARCH-in-mean models: 1 January 1986 through 1 September 1991*

	BF/DM	DG/DM	FF/DM	IL/DM	SP/DM
			3 Months		
Skewness	1.15	1.31	−0.21	−3.24	0.29
Kurtosis	4.90	7.10	4.05	16.58	6.50
BJ-test	25.45***	67.86***	3.66	650.69***	36.17***
KS-1	15.20***	19.52***	0.49	120.69***	0.99
KS-2	10.35***	48.34***	3.17*	530.29***	35.17***
LM(1)	0.79	0.95	0.01	1.14	0.72
LM(5)	7.81	2.94	3.88	8.40	27.37***
			6 Months		
Skewness	0.23	−0.66	−0.06	−0.17	0.60
Kurtosis	4.50	3.86	3.08	3.27	3.89
BJ-test	7.06**	7.16**	0.06	0.53	6.47***
KS-1	0.61	5.01**	0.04	0.32	4.19**
KS-2	6.46**	2.16	0.02	0.20	2.29
LM(1)	0.07	0.52	0.10	3.24	0.17
LM(5)	1.04	7.44	3.06	8.21	4.33
			12 Months		
Skewness	0.41	0.32	−0.59	0.00	0.47
Kurtosis	3.73	4.12	3.80	3.39	3.47
BJ-test	3.49	4.83*	5.89*	0.45	3.19
KS-1	1.92	1.18	4.07*	0.00	2.54
KS-2	1.53	3.64*	1.82	0.47	0.64
LM(1)	0.34	2.08	2.02	3.06	0.16
LM(5)	3.01	5.47	3.02	4.63	6.85

Notes: The BJ-test denotes the Bera–Jarque test for normality; KS-1 and KS-2 pertain to the Kiefer–Salmon normality test for skewness and kurtosis, respectively; * (**) [***] denotes rejection at the 10% (5%) [1%] level of the normality hypotheses.

Table 7. *Diagnostics for GARCH-in-mean models: 1 January 1986 through 1 September 1991*

	BF/DM	DG/DM	FF/DM	IL/DM	SP/DM
			3 Months		
Skewness	1.27	1.23	−0.16	−0.23	−0.21
Kurtosis	5.49	7.17	4.00	7.11	3.02
BJ-test	36.24***	67.55***	3.18	49.24***	0.50
KS-1	18.40***	17.45***	0.30	0.62	0.50
KS-2	17.83***	50.09***	2.88*	48.62***	0.00
LM(1)	0.82	1.04	0.11	0.94	1.16
LM(5)	7.68	2.91	5.47	0.96	5.18
			6 Months		
Skewness	0.20	−0.47	−0.07	−0.24	0.60
Kurtosis	2.84	3.97	3.09	3.47	3.89
BJ-test	0.54	5.20*	0.08	1.30	6.47**
KS-1	0.47	2.52	0.06	0.66	4.19**
KS-2	0.07	2.68	0.03	0.64	2.28
LM(1)	0.08	1.81	0.11	1.57	0.17
LM(5)	2.18	5.06	3.12	6.24	4.33
			12 Months		
Skewness	0.41	0.42	−0.43	−0.05	0.26
Kurtosis	3.73	4.95	3.01	2.80	3.06
BJ-test	3.46	12.97***	2.13	0.15	0.79
KS-1	1.92	1.98	2.12	0.03	0.77
KS-2	1.54	10.99***	0.01	0.11	0.01
LM(1)	0.34	1.35	0.95	0.79	0.26
LM(5)	3.01	4.58	2.70	4.75	6.37

Notes: The BJ-test denotes the Bera–Jarque test for normality; KS-1 and KS-2 pertain to the Kiefer–Salmon normality test for skewness and kurtosis, respectively; * (**) [***] denotes rejection at the 10% (5%) [1%] level of the normality hypotheses.

ARCH-in-mean models, rejection of the null hypothesis of no heteroscedasticity occurs in only one case (the SP/DM exchange rate), whereas the GARCH-in-mean models *all* result in statistically insignificant test statistics.

VI. CONCLUSIONS

This paper has examined exchange risk premia using survey data for a set of EMS exchange rates relative to the Deutschmark over the 1986–1991 period. The methodologies used in previous empirical research on premia in the pricing of forward foreign exchange usually involve measurement of time-varying risk premia conditional on market efficiency or rational expectations. We implemented an alternative approach to measure premia. The approach involves application of survey data to allow the direct measurement of risk premia from the forward discount decomposition into its two components – expected depreciation and the risk premium. We extended the analysis of Domowitz and Hakkio (1985) to model time-varying risk premia in the pricing of forward foreign exchange that do not require us to assume rationality on the part of economic agents. We find considerable support for the presence of time-varying risk premia in the pricing of forward foreign exchange. The estimated premium models – ARCH-in-mean and GARCH-in-mean – indicate that the time-varying premia can be explained by the conditional standard deviation of the expected rate of depreciation. In particular the GARCH-in-mean model appears to be reasonably succesful in accounting for both time-varying risk premia and conditional heteroscedasticity. Our findings basically contrast with the results of Domowitz and Hakkio (1985), who found only minimal support for the ARCH-in-mean model for some of the major currencies relative to the US Dollar in the period 1973-82.

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